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Trends in Prenatal Sex Selection and Girls' Nutritional Status in India

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Abstract

We study whether changes in prenatal sex selection across regions in India are associated with changes in girls' nutritional status. We exploit geographic variation in the incidence of prenatal sex selection and apply a triple difference approach comparing changes in the nutritional status of girls relative to boys across regions and over time. We find a reduction in girls' malnutrition in regions with an increasing incidence of prenatal sex selection. (JEL codes: J13, J16, I1, O12)

Keywords: son preference, prenatal sex selection, ultrasound, sex ratio at birth, gender discrimination, child health

1 Introduction

Son preference in India and other Asian countries has been found to be associated with differential parental investment in boys and girls, resulting in gender gaps in child health outcomes and mortality (e.g. Arnold et al. 1998; Deaton 1997; George 1997; Miller 1981; Kishor 1993; Rose 1999; Rosenzweig and Schultz 1982; Pande 2003; Barcellos et al. 2010). With the advent of prenatal sex determination technologies, parental behavior and thus girls' outcomes could be changing over time. On one hand, parents might respond to the technology by resorting to discriminatory investment against girls even prior to birth, possibly leading to deterioration of girls' births and postnatal outcomes. On the other hand, increased sex-selective abortions might imply that families who choose to give birth to girls might be less biased against girls, which should, on average, improve the outcomes of girls who are born.

In this article, we study the impact of prenatal sex selection on girls' nutritional outcomes. We first examine trends in the incidence of prenatal sex selection across regions in India. We show that sex ratios at birth have increased considerably in northwestern states relative to other regions and this increase can largely be attributed to the practice of prenatal sex selection. We then examine changes in nutritional outcomes of girls and boys in northwestern states relative to other states over time by estimating triple-differences models. This strategy allows us to account for time-varying factors at the state level that affect children's outcomes, in addition to control for state-fixed factors that differentially affect both genders and differential trends in boys' and girls' outcomes over time. Our results show that an increase in the practice of prenatal sex-selection in northwestern states is associated with a reduction in the prevalence of girls' malnutrition (relative to boys).

Our article is related to a limited number of recent studies (Shepherd 2008; Lin et al. 2009; Almond et al. 2010; Hu and Schlosser 2011) that examine the effects of prenatal sex selection on girls' outcomes. Lin et al. (2009) find a reduction in girls' mortality in Taiwan associated with the increasing use of prenatal sex selection. In contrast, Shepherd (2008) finds inconclusive evidence for a reduction in girls' mortality in India, whereas Almond et al. (2010) find an increase in girls' neonatal mortality following the introduction of ultrasound technology in China. Hu and Schlosser (2011) examine the impacts of prenatal sex selection in India on girls' nutritional status and mortality, and explore the possible channels linking between prenatal sex selection and girls' outcomes.[1]

We add to this limited number of studies in several aspects. First, compared to some other countries, India's high degree of heterogeneity in female discrimination and the differences in the practice of prenatal sex selection across regions makes it an interesting case for studying the effects of prenatal sex selection. This is also due to India's large variation in family sizes across households. Second, while previous studies have focused on mortality, we provide evidence on girls' nutritional status based on intermediate outcomes prevalent in a high proportion of Indian children, which have vital life-long consequences for human capital development and well-being (see, e.g. Currie 2009; Case and Paxson 2010). Third, our empirical approach, which is based on a triple-differences model, allows us to control for several confounding factors such as regional variation in son preference and differential trends across regions, providing a powerful way to reject alternative explanations for the observed results. Finally, compared to Hu and Schlosser (2011), who apply a continuous version of a triple-differences model using sex ratios at birth as the main explanatory variable, we abstract from the specific cohort measure of sex ratios and examine general trends across regions that differ in the prevalence of sex-selective abortion.

The rest of the article is organized as follows: In the next section, we review the literature and describe the institutional background of unbalanced sex ratios and prenatal sex selection in India. Section 3 describes

[1] Other related study is Bharadwaj and Nelson (2010) who examine gender differences in prenatal investment in countries with strong son preference, including India, although they do not estimate the impact of sex-selective abortion.

the data. Section 4 lays out the empirical strategy and presents the results. Finally, Section 5 concludes.

2 Background and Institutional Framework

India has long exhibited a female deficit dating back to the 19th century (see Visaria 1971; Miller 1981, 1984; Dyson and Moore 1983; Sen, 1992, 2003 who introduced the concept of "Missing Women"). Until the late 1970s, gender imbalances were mostly attributed to excess female mortality due to maltreatment and neglect and in extreme cases to female infanticide (see, e.g. Dreze and Sen 1997; Das Gupta 1987). Since the late 1980s, however, male to female ratios (MFRs) at birth have sharply increased, especially in northern and western states, which are regions historically known for strong son preference and gender discrimination. The increase in sex ratios at birth happened in concurrence with the spread of prenatal sex determination technologies, leading many to suggest that sex-selective abortion is likely to be a major contributing factor (see, e.g. Das Gupta and Bhat 1997; Arnold et al. 2002; Bhat 2002; Bhaskar and Gupta 2007; Retherford and Roy 2003).[2]

While abortion was legalized in India under the Medical Termination of Pregnancy Act (MTP) in 1972, sex determination during pregnancy was first made possible only in the late 1970s by the use of amniocentesis (Jefferey et al. 1984), and did not become widely accessible until the advent of ultrasound technologies in the 1980s.

Since the late 1980s, a continued decline in desired fertility coupled with a slower decline in the total number of desired sons increased the pressure to have sons at lower parities, thus raising the demand for prenatal sex selection (Das Gupta and Bhat 1997). At the same time, economic development and trade liberalization accelerated the supply of prenatal sex determination technologies. The diffusion process took place from urban to rural areas and from households of high socioeconomic status to those of low socioeconomic status (Khanna 1997).

In response to increasing public pressure from several NGOs and women's organizations to eliminate the practice of sex-selective abortion, the government of India passed the Prenatal Diagnostic Techniques

[2] Changes in the enumeration of girls or misreporting of age are unlikely to explain this upward trend since similar increases in sex ratios at birth are also observed among Indian populations living in the USA, Canada, and the UK where birth registration is nearly complete and accurate (see Abrevaya 2009; Almond and Edlund 2008; Almond et al. 2009; Dubuc and Coleman 2007). We provide in the next section further evidence that suggests that prenatal sex selection is the main cause for the upward trend in sex ratios at birth.

Regulations and Misuse Act (PNDT Act) in 1994, thus making it illegal to use ultrasound or amniocentesis in order to determine the sex of a fetus. However, this legislation proved to be ineffective and the practice of sex-selective abortion continued to spread (see e.g., George 2002; Kishwar 1995). In fact, as the publication of sex ratio figures from the 2001 Census revealed a continuing increase in the MFR at ages 0–6 years, the Indian government was further pressed to enforce and expand the legal power of the PNDT Act. Recent reports indicate some improvement in the enforcement of the act. However, sex-selective abortion is still being practiced extensively and enforcement of the law appears to be difficult, if not impossible, to achieve (see e.g., Subramanian and Selvaraj 2009; Portner 2010).

It has been estimated that about 0.48 million girls per year were selectively aborted in India during 1995–2005, which represents 6.2% of all potential female births (Bhalotra and Cochrane 2010). Estimates for northern and western regions are considerably higher. For example, Kulkarni (2007) estimates that out of 168 997 expected female births in Punjab in 2001, 19% (31 648) went missing.

3 Data

The data for our main empirical analysis are taken from the National Family Health Survey (NFHS). The NFHS is a large-scale, multi-round survey conducted in a nationally representative sample of households throughout India. Three rounds of the survey were conducted in 1992–1993, 1998–1999, and 2005–2006. Each round covered approximately 90 000 households that contained more than 500 000 individuals and was designed to provide state-level and national-level estimates. The survey includes detailed information on the demographic and socioeconomic background of the household members, as well as additional modules designed to investigate health, fertility, and mortality.

We pooled the three survey rounds and selected only households with ever-married mothers aged between 15 and 49 years in order to obtain a consistent sample across the three rounds.[3] Table 1 reports summary statistics for girls and boys included in our sample (columns 1 and 2), and differences by gender (column 3). The sample includes the youngest two children aged less than three of ever-married women sampled in one of the

[3] Only eligible individuals (ever-married women age 13–49 years in round 1, ever-married women age 15–49 years in round 2, and women age 15–49 years and men age 15–54 years in round 3) and their children were selected for the long-form questionnaires of the survey.

Table 1 Summary statistics

	Girls (1)	Boys (2)	Difference (3)
Urban	0.236	0.237	−0.002 (0.003)
Index of mass media exposure	0.799	0.824	−0.025 (0.010)
Wealth index	2.82	2.86	−0.041 (0.009)
Mother's age	25.2	25.3	−0.098 (0.024)
Mother's age at first birth	19.0	19.0	0.000 (0.025)
Mother's education			
No education	0.541	0.538	0.004 (0.004)
Primary school	0.158	0.148	0.009 (0.003)
Secondary school	0.247	0.256	−0.010 (0.005)
Higher	0.053	0.057	−0.004 (0.002)
Missing	0.001	0.001	0.000 (0.000)
Father's education			
No education	0.300	0.291	0.009 (0.003)
Primary school	0.193	0.189	0.004 (0.005)
Secondary school	0.381	0.389	−0.008 (0.006)
Higher	0.119	0.124	−0.005 (0.003)
Missing	0.006	0.007	−0.001 (0.001)
Religion			
Hindu	0.792	0.792	0.001 (0.002)
Muslim	0.158	0.156	0.002 (0.003)
Other religion	0.049	0.052	−0.003 (0.002)
Missing	0.001	0.001	0.000 (0.000)
Number of children in the family	2.920	2.927	−0.007 (0.009)
Mother wants another child	0.486	0.363	0.123 (0.009)
Sample size	36 940	39 560	76 500

Note: The table reports summary statistics for girls and boys (columns 1 and 2) included in the analysis samples and differences between the characteristics of girls and boys (column 3). Standard errors of the differences clustered at the state level are reported in parenthesis. The sample pools rounds 1, 2, and 3 of the NFHS and includes the two youngest children under 3 years of age of ever-married women with valid anthropometric data. Observations are weighted using national-level weights.

three NFHS rounds who have valid anthropometric data.[4] There are 76 500 children (36 940 girls and 39 560 boys) who satisfy these criteria. Household characteristics reported in the table are used as control variables in the empirical analysis.

[4] The first round of the NFHS collected anthropometric data for the youngest two children in the household who were under 4 years of age. The second round restricted the anthropometric data collection to the youngest two children in the household under the age of 3 years, whereas the third round extended the data collection to all children in the household under 5 years of age. To be consistent across survey rounds, we restrict the sample to the youngest two children under 3 years of age. In practice, our results are insensitive to these restrictions.

Most children (~75%) in the sample live in rural areas. About half of the children have mothers with no formal education and about 30% have fathers with no formal education. Mothers' ages at first birth are relatively low at 19 years, on average. At the time of the survey, mothers were 25 years old, on average, and they had an average of three children.[5] Column 3 shows that girls tend to be born into more disadvantaged families than boys. Their families have lower wealth levels, lower parental education, and a lower degree of exposure to mass media.[6] A possible explanation for differences in family characteristics of girls versus boys is the practice of sex-selective abortion. Another fact worth noting is that although girls do not appear to have more siblings than boys, probably due to the fact that a large proportion of children in the sample come from households with incomplete fertility, we see that mothers of girls are more likely to report that they want to have another child.

We use two alternative data sources for measuring the state-year MFR. The first source uses data on the number of males and females ever born based on the women's self-reported fertility history in the NFHS. To reduce measurement error, we compute MFR using the ratio of a 7-year moving average of the number of male births to female births by year and state using the pooled data of the three NFHS rounds. The second source of data comes from the population counts from the Indian decennial censuses. In this case, we compute MFR using information on the total number of children at age 0 by gender, state, and census year. MFR based on NFHS data has the advantage that it uses information from all children ever born so that it is not affected by differential mortality or living arrangements. However, it is susceptible to differential recall and sampling error. MFR based on census counts does not suffer from report or sampling errors, but includes only children who are alive at census date. We found that both measures are highly correlated.

4 Empirical Strategy and Results

We exploit variation in the incidence of prenatal sex selection across time and regions to examine its effect of on girls' nutritional outcomes.

5 Note that the average number of children in our samples does not represent completed fertility as most women are still in their fertile years.

6 The index for media exposure is defined by the sum of indicators for exposure to TV, radio, and newspapers or magazines. Each indicator receives a value of one if the mother reported exposure of at least once a week or almost every day. The wealth index is a constructed index provided in the NFHS data. The index is based on household assets and housing characteristics and denotes the wealth quintile of the household relative to all households sampled in the same survey round.

4.1 Incidence of Prenatal Sex Selection

One potential limitation of our analysis is that we do not directly observe the practice of sex-selective abortion. However, we do observe its consequences, primarily the abnormal sex ratio at birth. Figure 1 shows state variation in MFRs at age 0 as reported in census records from 1961 to 2001.[7] Until the 1980s, MFR at birth did not exceed the normal ranges of 103–107 males per 100 females found in various large-scale studies (e.g. Visaria 1971; Jacobsen et al. 1999). Increases in MFR at birth become evident at the transition points of 1981, 1991, and 2001, which overlaps with the diffusion of ultrasound technologies in India.

Interestingly, there are large variations in MFRs (both in their levels and in their growth rates) even across those states that appear to have a strong preference for boys. To illustrate this point, Table 2 reports MFR at birth by state for various census years (columns 1–5) and indicators of fertility and preference for number and sex composition of children based on tabulations from the first round of the NFHS (columns 6–9).[8] States are grouped by region.

The largest increases in sex ratios at birth are found in the northern and western states, which are characterized by a strong degree of son preference. In Punjab, for example, while MFR was within the normal range between 1961 and 1981, it increased dramatically between 1981 and 2001 from 106 to 129. In Gujarat, MFR remained at 103 between 1961 and 1981, but increased from 103 to 116 between 1981 and 2001. Both northern and western states appear to have strong son preference as manifested by the ideal sex ratio reported by mothers (1.46 and 1.29) and the proportion of mothers who desire a larger number of sons than daughters (0.49 and 0.38).

While strong son preference is found in states with upward trends in MFR, we also observe that in several states in the northeast, central, and eastern regions with similar strong preferences for sons there was either no increase in MFR or only a mild one. In Madhya Pradesh, for example, mothers reported an ideal sex ratio of 1.44, but sex ratios at birth

[7] MFR from census records number of children aged 0–11 months who are alive on census date. So, they also reflect gender differentials in infant mortality rates and living arrangements. In the next paragraphs we refer to MFR at age 0 as MFR at birth since changes in mortality and living arrangements by gender had only a negligible contribution to the overall trends in this variable.

[8] Women with living children were asked: 'If you could go back to the time you did not have any children and could choose exactly the number of children to have in your whole life, how many would that be?' Women with no living children were asked, 'If you could choose exactly the number of children to have in your whole life, how many would that be?' All women who gave a numerical response to the question on the ideal number of children were also asked how many of these children they would like to be boys, how many they would like to be girls, and for how many the sex would not matter.

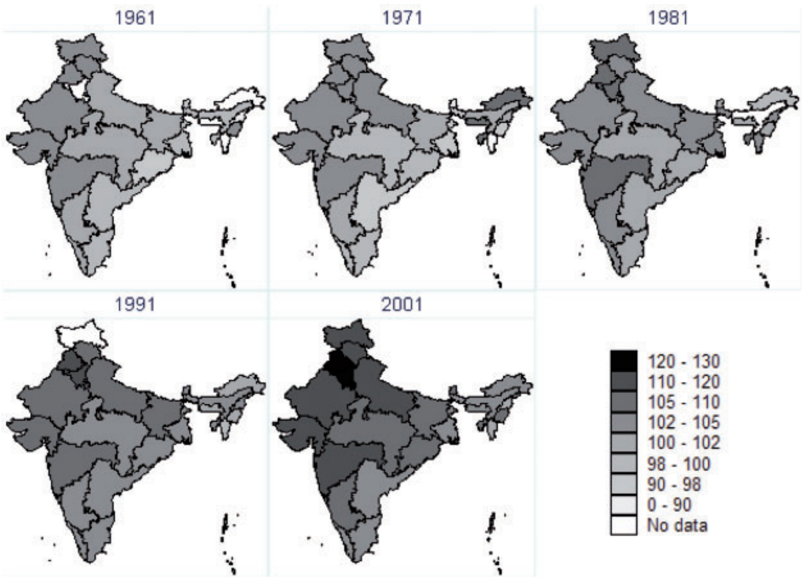

Figure 1 MFR at birth.

remained close to natural levels (MFR of 106 in 2001). Southern states are usually characterized by a low degree of son preference and stable sex ratios at birth.

Evidence reported in Table 2 suggests that strong son preference cannot alone explain state variation in MFR since there are several states with strong son preference that have not shown any significant increase in MFR. Indeed, a factor that distinguishes between states with similar son preference but different incidence in prenatal sex-selection is economic development. Table 3 shows that northern and western states are generally more economically developed than states in the northeast, central, and eastern regions. This is reflected by a higher wealth index, income per capita, share of households with electricity, and degree of exposure to mass media (TV). On the other hand, development and income levels in many of the northern and western states are comparable to those in southern states where sex ratios have remained balanced.[9]

Overall, evidence above suggests that the primary factors that characterize states with an increasing MFR are a strong preference for boys and

[9] In terms of women's educational level and religion, it is hard to find a clear pattern that differentiates states with increases in MFR from the rest. For example, MFR has increased significantly in Punjab, which has a high proportion of Sikhs and also in Himachal Pradesh where the majority of the population is Hindu.

Table 2 MFRs and fertility preferences by state

	MFR at age 0					Fertility preferences			
	1961	1971	1981	1991	2001	Number of children	Ideal number of children	Ideal MFR	Share who wants more sons than daughters
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
North	103	103	105	111	117	2.46	2.75	1.46	0.49
Delhi	104	105	105	110	117	2.36	2.52	1.25	0.30
Haryana	N/A	104	108	115	124	2.45	2.56	1.41	0.45
Himachal Pradesh	102	104	103	108	115	2.29	2.36	1.30	0.37
Jammu & Kashmir	102	103	107	N/A	114	2.58	2.77	1.48	0.49
Punjab	103	105	106	117	129	2.46	2.57	1.46	0.48
Rajasthan	103	102	104	108	112	2.49	3.02	1.55	0.58
West	103	103	104	108	113	2.23	2.56	1.29	0.38
Gujarat	103	103	103	109	116	2.24	2.60	1.33	0.42
Maharashtra	103	103	105	107	111	2.22	2.54	1.27	0.36
Northeast	98	102	102	104	104	2.73	3.33	1.33	0.40
Arunachal Pradesh	N/A	109	100	101	103	2.55	4.67	1.41	0.43
Assam	98	101	N/A	105	105	2.74	3.17	1.38	0.44
Manipur	102	94	101	102	106	2.89	3.74	1.36	0.43
Meghalaya	N/A	106	100	101	104	2.78	4.62	1.01	0.14
Mizoram	N/A	102	N/A	99	100	2.66	4.29	1.18	0.33
Nagaland	64	101	103	102	102	2.99	4.03	1.12	0.28
Tripura	99	106	106	103	105	2.43	2.57	1.28	0.33
Sikkim	95	88	101	105	106	2.32	2.23	1.13	0.22
Central	100	102	104	107	110	2.47	3.28	1.52	0.55
Madhya Pradesh	101	99	101	104	106	2.30	3.12	1.44	0.52
Uttar Pradesh	100	104	105	109	112	2.55	3.36	1.55	0.57
East	99	100	103	106	106	2.29	3.03	1.41	0.45
Bihar	101	102	104	108	107	2.38	3.40	1.56	0.56
Orissa	97	98	102	103	106	2.23	3.01	1.36	0.45
West Bengal	99	98	103	104	104	2.19	2.58	1.25	0.31
South	100	99	102	104	105	2.08	2.48	1.17	0.23
Andhra Pradesh	99	98	101	103	104	1.99	2.75	1.25	0.33
Goa	105	105	105	104	106	2.34	2.69	1.20	0.28
Karnataka	101	101	102	104	106	2.30	2.53	1.20	0.27
Kerala	101	99	102	104	103	2.07	2.62	1.12	0.18
Tamil Nadu	99	99	101	103	105	2.00	2.08	1.07	0.11

Note: Columns 1–5 report MFRs at age zero by state for various census years. Columns 6–9 report indicators for fertility, desired fertility, and son preferences based on mothers' reports in the first round of the NFHS. Tabulations for Sikkim are based on the second round of the NFHS as Sikkim was not sampled in the first round. Summary statistics reported in columns 6-9 are computed using state-level weights. N/A denotes data not available.

Table 3 State characteristics

	1990 NDP (Rs. Crore – 10 m. Rs)	1990 NDP PC (Rs.)	Household characteristics						Mother's characteristics		
			Urban	Wealth Index	HH with electricity	Religion			Average years of schooling	TV exposure	Illiterate
						Hindu	Muslim	Other			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
North			0.29	3.55	0.75	0.79	0.05	0.16	2.84	0.40	0.66
Delhi	10 243	11 057	0.92	4.79	0.96	0.82	0.10	0.08	6.35	0.83	0.37
Haryana	12 238	7508	0.26	3.90	0.86	0.89	0.04	0.07	3.01	0.49	0.64
Himachal Pradesh	2521	4910	0.10	3.62	0.92	0.97	0.01	0.02	3.62	0.47	0.50
Jammu & Kashmir	2908	3816	0.18	3.74	0.88	0.77	0.17	0.06	3.91	0.50	0.57
Punjab	16 738	8318	0.28	4.26	0.94	0.38	0.01	0.61	3.88	0.57	0.53
Rajasthan	18 281	4191	0.20	2.79	0.54	0.92	0.06	0.02	1.36	0.18	0.82
West			0.39	3.56	0.77	0.81	0.11	0.08	3.85	0.44	0.52
Gujarat	24 180	5891	0.35	3.60	0.78	0.89	0.09	0.02	3.61	0.39	0.55
Maharashtra	58 137	7439	0.42	3.54	0.76	0.76	0.13	0.11	3.97	0.47	0.50
Northeast			0.16	2.69	0.31	0.61	0.21	0.18	3.13	0.22	0.55
Arunachal Pradesh	460	5398	0.15	3.17	0.62	0.35	0.01	0.64	2.25	0.29	0.70
Assam	9498	4281	0.12	2.44	0.20	0.67	0.28	0.04	2.80	0.18	0.59
Manipur	723	3976	0.32	3.55	0.64	0.62	0.06	0.31	4.44	0.38	0.48
Meghalaya	767	4375	0.19	3.10	0.43	0.09	0.02	0.89	3.26	0.24	0.51
Mizoram	306	4474	0.49	3.82	0.76	0.02	0.00	0.98	5.69	0.25	0.08
Nagaland	579	4990	0.21	3.64	0.78	0.05	0.01	0.94	4.11	0.23	0.43
Tripura	917	3370	0.20	2.96	0.47	0.87	0.08	0.05	4.01	0.34	0.42
Sikkim	213	5302	0.14	3.73	0.80	0.60	0.01	0.38	3.72	0.56	0.49
Central			0.21	2.69	0.44	0.86	0.12	0.01	2.01	0.21	0.75

(continued)

Table 3 Continued

	1990 NDP (Rs. Crore – 10 m. Rs)	1990 NDP PC (Rs.)	Household characteristics						Mother's characteristics		
			Urban	Wealth Index	HH with electricity	Religion			Average years of schooling	TV exposure	Illiterate
						Hindu	Muslim	Other			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
Madhya Pradesh	26 515	4049	0.22	2.85	0.65	0.93	0.05	0.02	1.98	0.27	0.74
Uttar Pradesh	49 496	3590	0.20	2.62	0.34	0.83	0.16	0.01	2.03	0.19	0.76
East			0.19	2.46	0.24	0.83	0.15	0.02	2.40	0.21	0.66
Bihar	22 787	2660	0.15	2.32	0.17	0.82	0.16	0.02	1.78	0.13	0.78
Orissa	9664	3077	0.15	2.42	0.29	0.97	0.01	0.02	2.16	0.16	0.67
West Bengal	31 500	4673	0.27	2.67	0.30	0.76	0.22	0.02	3.30	0.33	0.51
South			0.31	3.39	0.65	0.82	0.11	0.07	3.72	0.43	0.54
Andhra Pradesh	29 867	4531	0.26	3.20	0.65	0.88	0.08	0.04	2.48	0.39	0.69
Goa	1024	8797	0.50	4.32	0.92	0.67	0.05	0.27	5.38	0.71	0.34
Karnataka	20 551	4598	0.33	3.27	0.66	0.86	0.11	0.03	3.13	0.40	0.61
Kerala	12 173	4200	0.28	3.89	0.61	0.54	0.26	0.19	6.76	0.42	0.16
Tamil Nadu	27 674	4983	0.35	3.42	0.66	0.88	0.06	0.06	4.07	0.50	0.50

Note: The table reports selected economic and demographic characteristics by state. Data on net domestic product (NDP) and net domestic product per capita (NDP PC) reported in columns 1 and 2 are based on reports of the Reserve Bank of India. Data on NDP are reported in crore (10 million rupees). Data on NDP PC are reported in Rupees. Means reported in columns 3–11 are based on tabulations from the first round of the NFHS. Means for Sikkim are based on the second round of the NFHS. Summary statistics reported in columns 3–11 are computed using state-level weights.

a higher degree of development and modernization (in combination). Still, there are some exceptions, such as the state of Rajasthan that is poorer and less developed but exhibits an increasing trend in MFR. There is also a clear geographical pattern that points to a higher incidence of prenatal sex selection in northern and western states. The fact that there are states with a strong preferences for boys and states with high levels of development that have not exhibited significant increases in MFR provide us with a heterogeneous group of states that are comparable to states with increasing MFR across different dimensions.

To examine the impact of prenatal sex selection on girls' nutritional status we exploit variation in the incidence of prenatal sex selection across regions over time by stratifying states in two groups: treated states (states in the northern and western regions) which, as discussed earlier, exhibit stronger son preference and an increasing trend in the practice of prenatal sex selection and comparison states (which include all other states). Specifically, we identify the following eight states as treated: Gujarat, Haryana, Himachal Pradesh, Maharashtra, Punjab, Rajasthan, Jammu and Kashmir, and Delhi. This list coincides with the classification of Bhat (2002) and adds the states of Jammu and Kashmir and Rajasthan to the classification proposed by Retherford and Roy (2003).

Figure 2 plots the MFR at birth by survey round in treated and comparison states for the cohorts of children born in the past 3 years prior to each survey date. The figure shows that MFR in treated states is higher than in comparison states. Moreover, while MFR increases sharply in treated states over the three survey rounds, MFR in comparison states appears to remain relatively stable.[10]

In Table 4 we examine whether the increase in MFR observed in treated states is directly related to the practice of sex-selective abortions by testing whether the propensity of giving birth to a boy is higher among families who might feel a stronger pressure to have a son and whether this propensity increased when prenatal sex determination became feasible. We report in this table the likelihood of a male birth at parity N (two or three) as a function of the sex composition of the older siblings who were alive at the time of conception using a linear probability model. We examine two samples: children born between 1975 and 1989 and children born from 1990 onwards. This split is meant to proxy for the availability of ultrasound technology. We assess the differential probability in treated states and in comparison states.

[10] We do not claim that prenatal sex selection is not practiced at all in other states, but rather that its effects are expected to be smaller relative to the treated group.

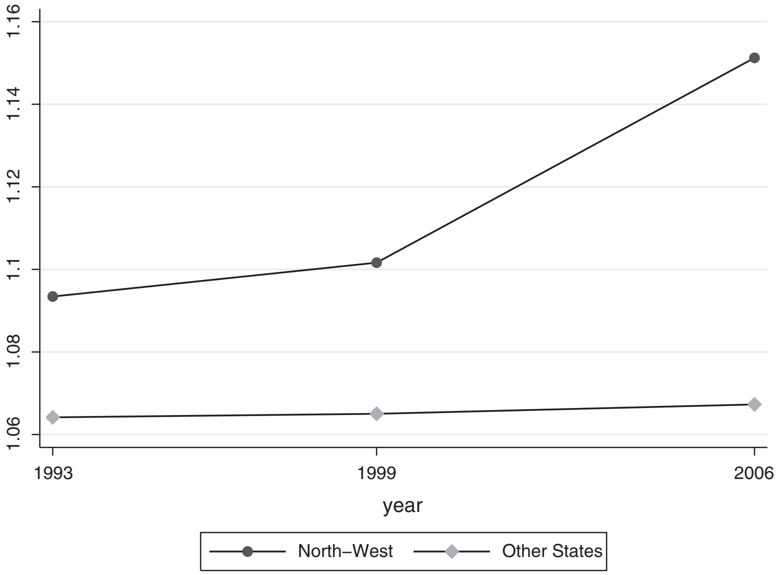

Figure 2 MFR at birth by survey round—children aged 0–35 months.
Notes: The figure plots average MFR at birth for the cohorts of children born within 3 years prior to each of the three survey rounds of the NFHS.

The results show that, prior to the 1990s, the likelihood of a male birth at parity two in treated states is not associated with the gender of the older sibling. At parity three, we already observe some positive association for households with two girls in treated states although the estimate is relatively small (2.7 percentage points). The differential probability of a male birth among households with one or two older girls increases considerably during the 1990s in treated states and is highly significant. Estimates for parity two show that households in treated states that have a girl at parity one, are almost 4 percentage points more likely to have a boy at parity two relative to households that have a boy at parity one. For parity three, we find that the likelihood of having a boy is 7.9 percentage points higher among households with two girls and 5.3 percentage points higher among households with a girl and a boy relative to households with two boys.

In sharp contrast to the pattern observed for treated states, we find no statistically significant differences in the likelihood of a male birth at parity two or three according to the sex composition of previous children among households in the other states during the post-ultrasound period.

Overall, the evidence presented above points to an increasing trend in the use of prenatal sex selection in the treated states as opposed to all remaining states, in which the practice of prenatal sex selection appears to be less common.

Table 4 Differential probability of a male birth at parity N as a function of sex composition of previous children: northern and western states versus other states

	Born between 1975 and 1989		Born after 1989	
	Treated states (1)	Other states (2)	Treated states (3)	Other states (4)
Panel A. Parity 2 (omitted category=Boy)				
Girl	0.004 (0.010)	−0.007 (0.003)	0.038 (0.008)	0.013 (0.009)
Sample size	*16 697*	*33 478*	*24 287*	*56 137*
Panel B. Parity 3 (omitted category=Boy-Boy)				
Girl–Girl	0.027 (0.006)	−0.020 (0.010)	0.079 (0.024)	0.012 (0.007)
Girl–Boy	0.003 (0.010)	0.011 (0.009)	0.053 (0.019)	0.007 (0.009)
Boy–Girl	0.017 (0.014)	−0.012 (0.010)	0.015 (0.014)	0.001 (0.011)
Sample size	*12 905*	*26 137*	*16 543*	*35 739*

Note: The table reports the differential probability of a male birth at parity 2 (panel A) and parity 3 (panel B) as a function of the sex composition of previous children. The table reports estimates for the subsample of treated states (columns 1 and 3) and all other states (columns 2 and 4). The sample includes all women aged 15–49 years surveyed in rounds 1–3 of the NFHS. Estimates reported in columns 1 and 2 are for children born between 1975 and 1989. Estimates reported in columns 3 and 4 are for children born in 1990 or afterwards. Regression estimates come from models that control also for twin status, mother's age, mother's education, mother's age a first birth, indicators for mother's religion, father's education, mother's mass media exposure, wealth, and rural/urban status. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis. Sample sizes are reported in italics.

4.2. Nutritional Status

We focus our analysis on children's nutritional status since it embeds information on various types of parental input and is measured more easily than most other outcomes and is therefore less likely to suffer from measurement error or recall bias. Children's nutritional status is assessed by anthropometric indicators based on height, weight, and age.[11] In particular, we consider three indicators of malnutrition: stunting, underweight, and wasting. All three are defined based on z-scores, which are computed by subtracting the median and dividing by the standard deviation of a reference population of the same age and gender. Specifically, a child is considered stunted if his or her height-for-age is at least two standard deviations below the median of the reference population (or the associated z-score is smaller than –2). An underweight child

[11] Height and weight were measured by the interviewers at the time of the survey using the United Nations guidelines (see Chapter 9 of IIPS 1995 for technical details).

has a weight-for-age at least two standard deviations below the median, and a wasted child has a weight-for-height at least two standard deviations below the median.

The three indicators capture malnutrition from different perspectives. Stunting reflects long-term malnutrition or cumulative nutrition from conception and is also affected by recurrent or chronic illnesses. Wasting measures acute malnutrition and represents the failure to receive adequate nutrition in the period immediately preceding the survey and may be the result of inadequate food intake or a recent episode of illness leading to weight loss. An important feature of the wasting indicator is that it does not depend on the accuracy of age reporting. On the other hand, it is more sensitive to seasonal shocks. Underweight is a composite index of chronic or acute malnutrition. Note that *z*-scores are normalized by gender and age so that they take into account that boys and girls may follow different growth trajectories. Our analysis uses *z*-scores based on the US National Center for Health Statistics (NCHS) standard, which was the most commonly used measure until 2006.[12] About 18% of the children aged 0–35 months included in our sample have missing values in at least one of the anthropometric indicators. Nevertheless, we do not find any significant gender differences in the likelihood of having a missing value in these indicators.[13] It is important to note that anthropometric indicators are only available for children who are alive at the survey date. We discuss the implications of this issue in the next section and assess the sensitivity of our results to the potential inclusion of anthropometric measures of dead children.

Figure 3 plots the nutritional outcomes for girls versus boys in treated and comparison states as measured in the first round of the survey, when the practice of prenatal sex-selection was still relatively low in most areas. Panels A through C plot the share of malnourished children (i.e. proportion underweight, stunted, or wasted) according to *z*-scores defined using the NCHS reference charts, panels D through F plot the equivalent indicators according to the 2006 WHO reference charts. Several remarks are in

12 A new international reference population was published by the World Health Organization in 2006. Our results are not sensitive to the specific reference chart used to define *z*-scores.

13 Height was not measured in the first round of the NFHS in Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu, and West Bengal because height measuring boards were not available at that time (IIPS 1995). Nevertheless, our results for weight are highly consistent with the results for height and our main results are unchanged when we limit the sample to states that have anthropometric data in the three survey rounds. We therefore believe that the lack of height data in round 1 for some states is unlikely to bias the main results.

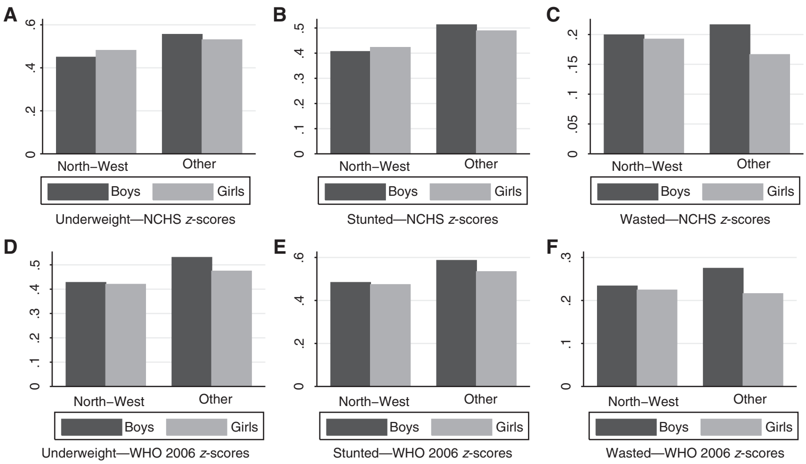

Figure 3 Gender gap in malnutrition by regions, 1992–1993.
Note: The figure plots the proportion of children who are underweight, stunted, and wasted by region (northwestern states vs. all other states) and gender. The sample includes children aged 0–35 months sampled in the first round of the NFHS (1993). Panels A through C plot malnutrition indicators according to the NCHS references charts. Panels D through F plot malnutrition indicators according to the WHO reference charts of 2006.

order. First, as noted in other studies (see, e.g. Rosenzweig and Schultz 1982; Sommerfelt and Arnold 1998; Griffiths et al. 2002), a comparison of nutritional outcomes between boys and girls is not very informative for inference of gender discrimination as these outcomes are also affected by biological processes, genetic endowments and illnesses that might differentially affect the genders.[14] Indeed, Figure 3 shows that in comparison states, boys are more likely to suffer from malnutrition than girls.[15]

[14] In Figure A1 we examine the proportion of boys and girls who are underweight using data from demographic and health surveys of various Sub-Saharan countries and India from the early 1990s. Interestingly, boys are more likely to be underweight than girls in most Sub-Saharan countries despite the fact that these countries do not show clear preferences for children sex composition.

[15] Other studies for India have also been inconclusive on whether there is a female disadvantage in nutritional outcomes of children when comparing outcomes by gender (e.g. Mishra et al. 1999; Pande 2003; Mishra et al. 2004; Tarozzi 2008; Barcellos et al. 2010).

Another important observation that emerges from this figure is that the gender gap in nutritional outcomes appears to be quite sensitive to the reference chart used for the standardization of z-scores. For example, the gender gap in the proportion underweight or stunted in treated states switches sign between panels 3A,B and 3D,E. A consistent pattern emerging from this figure is that girls' relative nutritional status in treated states is always lower than in comparison states. This is evident even in cases where girls do not appear to have a nutritional disadvantage relative to boys (as in panels c, d, e, and f).

Evidence presented in this figure suggest that the gender gap in nutritional status *per se* cannot be used as a direct indicator for gender discrimination, however, the comparison of gender gap across regions can be informative about the relative status of girls. An alternative approach for inference about girls' differential treatment would be to compare girls' nutritional status across regions. However, this comparison would need to account for all other differences between states that could affect girls' outcomes. To account for these differences, we examine changes in the outcomes of girls versus boys over time comparing between regions with an increasing prevalence of prenatal sex selection and regions where prenatal sex selection has remained relatively stable.

4.3. Effect of prenatal sex selection on girls' nutritional status

We analyze the change in girls' nutritional status in northern and western states relative to boys and relative to other states over the three survey rounds by estimating the following equation:

$$\begin{aligned} y_{is\tau} = &\alpha_{s0} + \alpha_{s1} female_i + \delta_{\tau 0} + \delta_{\tau 1} female_i + x_i'\beta + \gamma_{\tau 0} Treated_s \\ &+ \gamma_{\tau 1}(Treated_s * female_i) + \varepsilon_{is\tau} \end{aligned} \tag{1}$$

where $y_{is\tau}$ is the outcome of child i in state s and in survey round τ , α_{s0} and α_{s1} are vectors of gender-specific state fixed effects, $\delta_{\tau 0}$ and $\delta_{\tau 1}$ are vectors of gender-specific survey-round fixed effects, x_i is a vector of individual characteristics that include year of birth fixed-effects, mother's age at first birth, indicators for twin birth, residence in an urban area, religion, mother's and father's level of education, mother's age (grouped), wealth quintiles and mass media exposure. 'Treated' is an indicator that equals 1 if child i was born in a state with a high incidence of prenatal sex selection and 0 otherwise. The parameters of interest are $\gamma_{\tau 1}(\tau = 1998; 2005)$ which denote the differential change in girls' outcomes between 1992 (the year of the first round of the NFHS) and 1998 or 2005 (the years of the second and third rounds of the NFHS survey, respectively) in states with an upward trend in prenatal sex selection relative to states in which prenatal sex selection is rare and has not increased over time.

This approach is a *triple-difference* estimation strategy.[16] This strategy has the advantage that it allows us to control for state-level fixed factors that differentially affect boys and girls (e.g. the degree of discrimination against girls in a state). We also control for differential trends in boys' and girls' outcomes at the national level in a very flexible way by including gender-specific survey-round fixed effects. In addition, we control for time-varying factors that affect boys and girls similarly in treated and comparison states (e.g. improvement in access to health facilities). In this regard, it is important to note that we always include in our model the interactions between survey-round and treated, which gives us an indicator of the effect of unobserved time varying factors that differentially affect boys' outcomes in treated and in comparison states.

Table 5 reports the estimates from our triple-difference equation for a linear probability model of likelihood of being underweight, wasted, or stunted. The coefficients on the triple interaction terms, *female*round2*treated* and *female*round3*treated*, are negative for the three nutritional status indicators. The magnitude of the estimates is larger (in absolute terms) in the third survey round relative to the second round (see columns 4 and 6), which is consistent with the upward trend in MFR over the three survey rounds observed in Figure 2. For example, estimates for underweight (first row) suggest that girls' likelihood of being underweight decreased about 4 percentage points more in states with a higher incidence of prenatal sex selection relative to other states between the first and the second round of the NFHS. The reduction observed in the third round relative to the first is about 6 percentage points. The differential decrease is about 9 and 13%, respectively, of the outcome mean. Note that sex ratios at birth increased in treated states from 1.093 to 1.151 between the first and the third round, whereas in other states, they increased only slightly from 1.064 to 1.067. Estimates for the additional outcomes provide a similar picture of a decline in the proportion of girls who are wasted or stunted in treated states over the survey rounds.

Interestingly, with the exception of stunted, the coefficients for the two interaction terms, *round2*treated* and *round3*treated*, are generally small, and insignificant. This finding suggests that at least for underweight and wasting, there does not appear to be major changes over time in the status

[16] An alternative strategy is to replace the interactions between treatment indicator, survey rounds and gender with a main MFR effect and its interaction with female. This alternative strategy exploits the full variation in the extent of prenatal sex selection across time and regions. However, it requires a stronger assumption for identification. In Hu and Schlosser (2011) we use this alternative strategy and show that the two approaches yield very similar results.

Table 5 Effects on nutritional status of children by region and survey round

Outcome	Sample size (1)	Outcome mean (2)	Round 2 x treated (3)	Female x round 2 x treated (4)	Round 3 x treated (5)	Female x round 3 x treated (6)
Underweight	76 314	0.485	0.034 (0.040)	−0.043 (0.019)	0.000 (0.040)	−0.064 (0.025)
Wasted	69 784	0.179	0.004 (0.023)	−0.022 (0.017)	−0.009 (0.033)	−0.056 (0.016)
Stunted	69 571	0.433	0.037 (0.032)	−0.018 (0.020)	0.032 (0.040)	−0.034 (0.024)

Notes: Column 2 reports means of the dependent variables. Columns 3–6 report estimates from a triple-differences model that compares changes in nutritional outcomes of girls versus boys in treated versus comparison states over the second and third survey round relative to the first survey round. The treated group includes the following states: Gujarat, Haryana, Himachal Pradesh, Maharashtra, Punjab, Jammu and Kashmir, Rajasthan, and Delhi. The models control for state fixed effects and survey round indicators interacted with gender. In addition the model controls for twin status, mother's age, mother's age at first birth, and indicators for mother's religion, mother's education, father's education, mother's mass media exposure index, wealth index, and rural/urban status. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis.

of boys in treated states relative to other states, which to some extent alleviates the concerns that our results are confounded by omitted time-varying factors that affected child nutritional status in treated states.

As noted above, the anthropometric data is available only for surviving children. If prenatal sex selection has any impact on mortality, our analysis of nutrition will be based on a selective sample. For example, if increases in MFR are associated with a reduction in female child mortality, they might lead to an increase in the proportion of girls who are close to a survival threshold, thus attenuating the estimated effect of MFR on nutritional status. However, results from Hu and Schlosser (2011) suggest the effects on mortality are negligible or not significant, which mitigates the concern about selectivity issues in the nutritional status results.

We further assess the sensitivity of our results to differential mortality by imputing anthropometric measures for dead children. Specifically, we assign to dead children the lowest quintile or decile of outcomes observed for children who are alive and of the same age, gender, state, and survey round. The estimates based on the expanded sample, reported in Table A1, are very similar to our main results.

Taken together, our findings suggest that girls' nutritional status improved to a larger extent in states with an increasing trend in the use of prenatal sex selection.

5. Summary

In this article, we study the impact of prenatal sex selection on the nutritional outcomes of girls in India. We use data from the NFHS and exploit variation in prevalence of sex-selective abortion across regions and over time. More specifically, we examine changes in the outcomes of girls versus boys over the three survey rounds in northwestern states, which have experienced an upward trend in the practice of prenatal sex selection, relative to other states where the practice of prenatal sex selection has been relatively rare.

Our results show that states with an increasing trend in the practice of prenatal sex selection experienced a larger reduction in the proportion of malnourished girls. We show that our results cannot be explained by changes in mortality.

One important question is what are the possible factors that could explain the relative improvement in girls' nutritional status in states with a higher prevalence of sex-selective abortion? Besides the aforementioned selection channel through which girls are born into families with weaker son preference, girls could also be born into smaller families as prenatal sex selection allows parents reduce their reliance on son-based fertility stopping rules. Girls' outcomes might also be affected if families that selectively abort girls have different characteristics than those that do not practice sex-selective abortion. Finally, as girls become increasingly scarce and thus more valuable in the labor and marriage markets, parental incentives to invest in girls might increase even among families who do not practice prenatal sex selection. In Hu and Schlosser (2011), we further explore these underlying mechanisms linking prenatal sex selection and girls' nutritional outcomes. A better understanding of the mechanisms would also provide important information regarding the potential unintended consequences of banning sex-selective abortion and encourage alternative polices for promoting parental investment in girls.

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Appendix

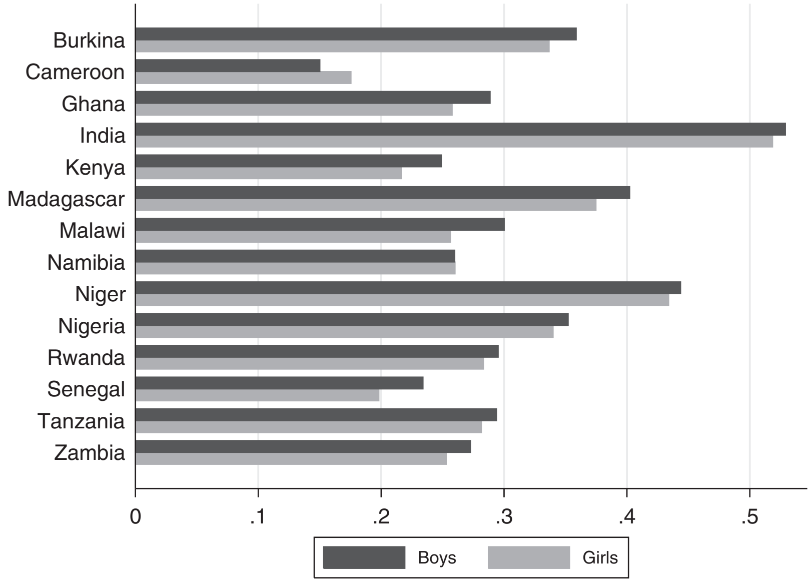

Figure A1 Proportion underweight by gender: children aged 0–35 months.
Note: The figure plots the proportion of children who are underweight in various countries. Data are based on the demographic and health surveys from the following years: Burkina 1993, Cameroon 1991, Ghana 1993, India 1992, Kenya 1993, Madagascar 1992, Malawi 1992, Namibia 1992, Niger 1992, Nigeria 1990, Rwanda 1992, Senegal 1992, Tanzania 1991, Zambia 1992. Data is weighted using national weights.

Table A1 Effects on nutritional status of children using imputing data for dead children

Outcome	Sample size (1)	Outcome mean (2)	Round 2 x treated (3)	Female x round 2 x treated (4)	Round 3 x treated (5)	Female x round 3 x treated (6)
Underweight						
Imputing with lowest quintile	82 437	0.522	0.022 (0.038)	−0.038 (0.017)	−0.008 (0.038)	−0.055 (0.024)
Imputing with lowest decile	82 437	0.525	0.023 (0.037)	−0.036 (0.014)	−0.007 (0.038)	−0.054 (0.023)
Wasted						
Imputing with lowest quintile	75 299	0.190	0.010 (0.030)	−0.022 (0.028)	−0.003 (0.039)	−0.068 (0.023)
Imputing with lowest decile	75 299	0.231	0.014 (0.037)	−0.023 (0.019)	0.010 (0.042)	−0.069 (0.025)
Stunted						
Imputing with lowest quintile	75 086	0.468	0.032 (0.034)	−0.015 (0.022)	0.019 (0.046)	−0.020 (0.026)
Imputing with lowest decile	75 086	0.476	0.032 (0.030)	−0.019 (0.017)	0.026 (0.040)	−0.030 (0.024)

Notes: The table reports triple differences estimates for the same model estimated in Table 5. The sample includes also children who were dead at the survey date. Anthropometric measures for dead children were imputed using data from the lowest quartile or decile of children of the same gender, age, state, and survey round. Column 2 reports means of the dependent variables. Observations are weighted using national-level weights. Standard errors clustered at the state level are reported in parenthesis.